Employability of young Italian males

after a jobless period, 1989-1998

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Abstract

In this paper we investigate the existence of negative jobless duration dependence and the impact of jobless spells on future wages. Our findings are relatively out of line compared to analogous explorations. We find evidence of very long unemployment duration of the young male labor force, much higher than reported anywhere else in Western Europe. Despite our findings on unemployment duration, negative unemployment duration dependence is modest. While the probability of re-employment decreases also in Italy as elapsed joblessness becomes longer, such decay is small. Finally, we show that young Italian males experiencing jobless periods in their early careers face small re-employment wage losses. Such losses do increase with the duration of joblessness, but here, once again, they are lower than reported in Unites States, Canada, UK, France and Spain.

Keywords: joblessness, duration dependence, long term unemployment, two stage least squares, selection problems. **JEL-codes:** J64, J24

1. Introduction

The aim of this paper is to answer to the following question: ”do youth jobless spells impact on employability and future wages in Italy?”.

The problem of European unemployment has spurred an immense amount of academic literature. In the 1970s the European unemployment rate started to rise sharply. From the mid-1980s until 1990, it gradually fell although it remained still higher than the early 1970s. Large differences in unemployment trends were observed across European countries. The European unemployment problem is magnified when focussed on youth workers: youth unemployment ranges from 18-23% in France, Italy and Spain in the late 1990s (Howell et all., 2006). The incidence of youth unemployment has been related to the effectiveness of the educational system at easing the transition on from school to work (see, for instance, OECD, 2000), to labour market institutions (such as unemployment benefits for the young, minimum wages, labour cost and flexibility, etc.), to the role of the family at providing income support (Bentolina and Ichino, 2000), and to the evolution of the relative size of the youth population (Korenman and Neumark, 2000).

Youth unemployment raises special concern in regard to its detrimental effects on long term welfare. A number of empirical studies of both U.S. and European countries suggest that the employability of jobless persons deteriorates as their joblessness persists. Blanchard and Summers (1989), Layard and Nickell (1987), Machin and Manning (1999), all emphasize the adverse effect of long term unemployment on human capital, skill obsolescence and stigma. The negative relationship between the duration of joblessness and the probability of being rehired persists also when selection issues related to workers’ heterogeneity are included in the analysis (Van den Berg and Van Ours, 1994 and 1996). Torelli and Trivellato (1989) study youth unemployment duration in Italy, confirming state dependence. Ordine (1992) also reports evidence of state and duration dependence for prime age Italian unemployed and first-job seekers. Recently, Addison, Centeno and Portugal (2004) offer new evidence that the hazard function exhibits strong negative dependence in the EU15.

Workers experience wage cuts after long unemployment spells. These could be related to reservation wages guided by low unemployment benefits or due to the depreciation of general and specific human capital skills during unemployment. Farber and Gibbins (1996) focus on the signaling problem associated with job loss: displacement may be particularly costly if it is used by prospective and future employers as a signal of bad worker performance. Previous literature on the United State (Ruhm, 1991; Addison and Portugal, 1989; Topel, 1990; Jacobson et al, 1993; Farber, 1993 and 1997; Seninger, 1997) and Canada (Houle and Van Audenrode, 1995) indicate that joblessness leads to wage losses of an order of magnitude of 10% to 15%. Kletzer and Fairlie (2003) find that the earnings loss of young workers after re-employment are also substantial, although smaller and less persistent than those observed for older and more established workers. Wage losses in EU could be distorted by wage rigidities, but evidence is still insufficiently conclusive and new studies are necessary. Among the existing contributions[1], Gregory and Jukes (1997), Cohen, Lefranc and Saint-Paul (1997) and Rosolia and Saint-Paul (1998) analyse the impact of joblessness on re-employment wages using respectively British, French and Spanish data: the wage loss of displaced workers in France and UK is similar to the one observed in US, while it is higher in Spain. On the other hand, Addison, Centeno, and Portugal (2008) report that there is scant evidence of a decline in reservation wages with longer joblessness in the EU15. Contini and Villosio (2006) deliver, instead, results in line with the weak evidence of unemployment duration on re-employment wages found here: using microdata from the same source as this paper, they suggest that prolonged unemployment spells in Italy have a modest negative impact on the wage growth of white-collar employees (up to 2.5 p.p. for a six-month spell), very slight on the blue-collars'.

Our focus is on the Italian labour market. The period of analysis is the late 1980s and the early 1990s. As explained before, the labour market was characterized by high and persistent youth unemployment, while the traditionally rigid labour market institutions were

[1] Other interesting studies are Lefranc (2003) Ahn and Garcia-Perez (2002) and Garcia-Perez and Rebollo-Sanz (2006)

undergoing change. A reform aimed at introducing flexibility was enacted in 1997 (Pacchetto Treu); its main innovation is the "contratto di collaborazione continuativo" ("co.co.co"), a mixed form of dependent work and self-employment, that delivered considerable flexibility and cost savings to the employers, and much less security to the employees.

Moreover the 1989-1998 period is a particularly interesting time window as it contains an entire business cycle, from the boom of the late 1980s to the deep recession of the early 1990s. We analyze jobless duration and post joblessness wage loss in Italy from October 1989 to December 1995, focusing on youth males who have had a relatively significant work experience and job attachment.

Two issues are faced in this exploration. First, we investigate the existence of negative jobless duration dependence. This provides the basis for a better understanding of the mechanisms that produce employer stigmatization, discouragement, and human capital depreciation over the course of joblessness. Second, we investigate the impact of jobless duration on wages, addressing the issues of self selection into employment and the endogeneity of jobless duration. The analysis provides preliminary signals of the jobless effects on the welfare of the young people in the longer term. Empirical analysis is based on WHIP (Work Histories Italian Panel), a unique employer-employee linked panel database which represents a very detailed source of information on the working histories of Italian employees.

According to a recent paper by Tatsiramos (2009), the average Italian unemployment duration for no-benefit recipients in the Nineties was about 12 months: 55.3 per cent of non-recipients in Italy were reported unemployed after 6 months, and 35.2% were still unemployed after 12 months. His estimates are based on EU-SILC data.

In this paper we provide new and more dramatic evidence on very long unemployment duration of Italy's young male labor force: in the course of the Nineties 40% of the unemployed had yet to find a job 36 months after entry in the jobless state. This share was down to 33% for those aged 16-19 on their first job, and reached 51% for the late starters (over 26). These numbers imply much longer unemployment duration than reported anywhere else in Western Europe. In addition, our results confirm the existence of negative duration dependence of young Italian males, although to a much weaker extent than all the above mentioned contributions.

The paper's contribution to the existing literature is threefold. Firstly, as already mentioned, we provide new and more dramatic evidence on very long unemployment duration of Italy's male labor force, much longer than reported anywhere else in Western Europe. Secondly, our results confirm the existence of negative duration dependence of young Italian males, but to a weaker extent than other known contributions. Thirdly, we offer new econometric evidence on the adverse effects of a period of joblessness on young Italian workers' wages upon re-employment. Also on this score, our results indicate that youth experiencing jobless periods in their early careers face significant, but small re-employment wage losses. Such losses do increase with the duration of joblessness, but, in line with other reported differences, they are much lower than those observed in Unites States, Canada, UK, France and Spain.

These findings are novel and relevant for the design of employment policy and for a better understanding of the effects of the European unemployment problem on the future welfare of young workers. Low unemployment duration dependence and modest job losses at re-employment may be good news, but they are overshadowed by the bad news on the length of unemployment. Here is where the real problem lies: perhaps not a great discovery, but certainly important to re-address policy instruments aimed at easing the tensions that run through the Italian labor market.

The paper proceeds as follows. Section 2 provides detailed information on the data. In section 3, we illustrate the model used in subsequent paragraphs. In section 4, we present the main empirical findings. Finally, section 5 briefly concludes.

2. Definitions and data

We use the Work Histories Italian Panel (WHIP), an employer-employee linked panel database developed by Italian Social Security administrative sources. The WHIP data[2] are a representative sample of the population of employees of the private sector (agriculture excluded), apprentices, self-employed, atypical (non-standard) contracts, as well as all public employees working without tenure (nowadays almost 50% of the young public employees) [3]. The sample - population ratio is 1:90. WHIP observations start in 1986 and, as of today, end in 2003. For reasons explained below we limit analysis to the 1988-1998 decade. The Italian Social Security Administration (INPS) collects data both on individual employees and firms for institutional purposes. The reference population is made up of the all individuals – Italian and foreign – who have been regularly employed or self-employed. Firms pay social security contributions to INPS for all workers on payroll. Data are recorded on a monthly basis, and therefore working spells can be precisely reconstructed. The WHIP database contains information on worker age, professional category, industrial sector, length of employment spells, geographical location, working contract and monthly wages[4]. In administrative archives, information not related to the specific interest of the Italian Social Security Administration (i.e. marriage status, children, etc.) is not present. On the other hand, the coverage and accuracy of administrative archives cannot be matched in any other dataset. Neither out-of-the-labor-force people, nor the unemployed who are not eligible for unemployment benefits appear in the INPS files.

We use data from January 1985 to December 1998. While more recent data are available, we focus on this period as it covers an observation period centered around years immediately preceding and following the 1993-94 recession. The Treu Reform of 1997 introduced new atypical contracts ("co.co.co", a mix of dependent work and self-employment which allowed great flexibility): while the "co.co.co" working spells are reported in the WHIP

[2] The WHIP is public use database. See http://www.laboratoriorevelli.it/.

[3] Social security contributions of the public sector are paid to a different institution (INPDAP). In principle INPDAP ought to have the same data as Italian Social Security Administration, but, unfortunately, the INPDAP data warehouse has never been completed and precise data on public employment are non-existent. The (unofficial) Italian count is based on Labour Force Survey data.

[4] Monthly wage is computed using information about yearly wage and the number of days worked during the year. The yearly wage is divided by the number of days and multiplied by 26 days. Real monthly wage is computed in the standard way.

database, earnings data are not available. As a consequence we exclude them from investigation as they would considerably fuzzy the picture that we obtain. The loss of information, however, is small as the utilization of the new contract became important only toward the end of 1998.

Exits from the databases reflect transitions from employment to non-employment or unemployment. In principle, we do not have any attrition problems because, once a certain group of individuals is selected, it is possible to follow them over the entire working life. There are, however, few exceptions:(i) movements from private to public activities go unrecorded if the job switch is accompanied by full tenure in the public sector (a very rare event for young people in the last twenty years) [5]; (ii) shifts from employment to school go also unrecorded. If they are temporary, they will be followed by re-entry in the labor market some time thereafter, in which case no information is lost. The case of no re-entry is the exception. Exceptions (i) and (ii) involve a small number of individuals. Instead exception (iii) - moves from regular employment into the parallel, hidden economy – are presumably frequent especially among individuals with certain characteristics – possibly, low educational attainment – but they are undetectable by definition.

As in other studies on the employability of young displaced workers (Topel and Ward, 1992; Kletzer and Fairlie, 2003) we restrict attention to individuals at the very beginning of their career, exhibiting a relatively high attachment to the labor market. In order to meet these requirements, our sample includes young male employees (aged 16-30) with the following characteristics:

- absent from the WHIP database in the 1985-88 time window: this insures that the job observed starting in 1989 is one's first working experience;
- first working experience between January 1989 and December 1993;
- employed for a period of 9 up to 23 months;

[5] This exclusion includes also to the military and police services. On the other hand individuals in draft age, employed by the same firm before and after the ten-month period corresponding to mandatory military service are considered as regularly employed throughout. Usually individuals complete military service before entering the labour market. Military service was mandatory until 1992, and then become voluntary. The Ministry of Defense estimates about 20 thousand /year new entries, about 0.5% of the Italian male population 20-29 and less than 3% of the total number of young entries in the labor market.

- started a jobless period between October 1989 and December 1995.

The reason for excluding individuals with working spells longer than 23 months is twofold: (i) we wish to observe working histories for 36 months after involuntary separation; (ii) only doing so can we exclude all recipients of unemployment benefits: we would, otherwise, risk to observe some and leaving out others, with no possibility to control. [6]

The idea of restricting attention to the "relatively attached" workers (those with initial employment spells at least equal to 9 months) allows to make our results comparable with other studies [7] Such a selection may introduce an "optimistic" bias in our results. On the other hand, the exclusion of those who had a working spell longer than 23 months may also introduce selection in the opposite direction, as many of these could be "very good" workers.

In addition, analysis ought to be restricted to individuals who have experienced involuntary separations. In order to single out involuntary separations, we exclude all the job-to-job movements and the job switches that take place within the same month. Thus, in first instance, we consider as involuntary all separations that last at least one month.[8] But we also test our results using a stricter definition of involuntary separation (lasting at least three months), obtaining only minor changes in the results (see par. 5).

[6] Complete information on "unemployment beneficiaries" is not available in WHIP. Unemployment benefits were seldom granted in the Nineties, with the exception of the construction sector. Workers were instead eligible for tor temporary layoff benefits (C.I.G.S. Cassa Integazione Guadagni Straordinaria), when unemployment exceeded 24months. CIGS payments are, instead, recorded in WHIP, involving about 3% of the sample individuals.

[7] In order to define labor market attachment, other authors have used tenures slightly longer than 9 months (e.g., by Topel and Ward, 1992; Kletzer and Fairlie, 2003). In our view, nine months is a reasonable long period to signal a sufficient degree of attachment (while, the risk of capturing short-term employment before a return to formal schooling or training is negligible).

[8] The difference between quits (voluntary separations) and layoffs (involuntary separations) is always very problematic in applied economics. Layoffs carry a stigma that may often lead interviewed people not to report them in surveys, or to instead report a quit. Administrative data may be more reliable if the payment of unemployment benefits is individually reported (as already explained, in our sample few workers are eligible for such provisions). In LFS-type surveys employers may be unwilling to report correctly: for example, in the early 2000 Fiat dismissed over 20 thousand workers, 25% declared layoffs and 75% quits. But all the reported quits were "forced resignations" accompanied by a generous severance payment made conditional to the acceptance of dismissal (reported by national press)..

Our final sample includes 2318 workers whose working history is followed for 36 months.[9] The average elapsed period of joblessness is about 20 months (and about 10 months for the sub-sample of individuals re-employed by the end of the period of study). For more details on sample composition see Table 1.

3. The Model

3. 1 Transition out of joblessness

We analyze the duration of joblessness with a view of investigating the existence of negative duration dependence and of understanding which factors influence the transition path. Negative duration dependence may be produced by declining job offer arrival rates, increasing reservation wages, or/and from an adversely shifting wage offer distribution (Addison et al. 2004). An alternative, while not exclusive, explanation suggests that negative dependence may be a consequence of sorting of the more employable among the jobless workers. We use a discrete-time hazard rate model (i.e. Narendranathan and Nickell, 1989; Jenkins and Garcia-Serrano, 2000). In particular, we consider all individuals from the moment they become jobless and are likely to exit thereafter. The probability of person i of being re-employed after t months, given that he has been jobless for t-1, is assumed here to be a standard logit hazard function:

$$h_{it} = \exp[\mathbf{x}_{it}'\boldsymbol{\beta}+\gamma(t)]/(1+\exp[\mathbf{x}_{it}'\boldsymbol{\beta}+\gamma(t)]) \quad (1)$$

where $\mathbf{x}_{it}$ is the vector of (time-constant and time-varying) covariates, $\boldsymbol{\beta}$ is a vector of parameters to be estimated and $\gamma(t)$ is some functional form of how the duration of the spell affects the hazard rate (baseline function). For the latter, we initially use a linear log-time

[9] Of the initial sample of 3199 individuals, about 7% are re-employed with atypical contracts and 16% become self-employed. Therefore, our sample is representative of the population of young males experiencing jobless periods and currently looking for jobs. We disregard, for the time being, the potential impact of multiple spells.

specification. In second instance, we use a flexible specification (duration-interval dummies) in order to avoid the potential estimation bias due to the specific assumption of the form of the baseline function (Meyer, 1990). Estimation of the model parameters can be done using standard software applied to a re-organized data set in which, for each person, there are as many data rows as there are time intervals at risk of the event (Allison, 1982; Jenkins, 1995; Jenkins and Garcia-Serrano, 2000). Since individuals might differ in unobserved terms like ability, effort, and taste and these differences could remain constant over time, we include unobserved heterogeneity in the specification of the hazard rate.

3.2 Post-joblessness wage analysis

The impact of joblessness on earnings of individuals who re-enter the job market is estimated on the following specification:

$$(2) \qquad \log(w_a) = \mathbf{z_a}\, \boldsymbol{\gamma_a} + \alpha_a \log(t) + \beta_a \log(w_p) + u_a$$

where $\log(w_a)$ is the logarithm of the post-joblessness real monthly wage, $\mathbf{z}_a$ is the vector of the explanatory variables that influence the post-joblessness wages but not the pre-joblessness earnings (i.e. attributes of the new job, changes of industry, working area and occupation, and actual local unemployment rates), $\log(t)$ is the logarithm of the elapsed joblessness duration (in months) and $\log(w_p)$ is the logarithm of previous job earnings. The estimation of the above equation raises two main econometric issues: selection and endogeneity.

Selection problems arise as a considerable fraction of the jobless individuals were not re-employed as of December 1998. For such individuals the effects of the determinants of post-joblessness earnings could be systematically different from those of re-employed people. The conventional two-step selectivity adjustment procedure proposed by Heckman (1979) is implemented, leading to consistent parameter estimates. We also use the suggestion of Hill, Adkins, Bender (2003) to calculate correct standard errors in presence of heteroskedasticity[10]. Heckman-type selection models normally need at least one "extra" explanatory factor that influences selection but not the subsequent outcome of interest. Often such "extra" exogenous variable is not available. This is our case: the same elements affecting selection will also affect the length of the unemployment spells (although some may not affect post-jobless wages). The Heckman selection models are also estimable without the extra variable, with results resting only upon the distributional assumptions of the residuals rather than upon the variation in the explanatory variable (Sartori 2003; Liao 1995). Despite the concerns expressed by Dufour and Hsiao (2006) on violations of the distributional assumptions, model identification based solely on distributional assumptions appears reasonable whenever valid extra variables are not available.

The endogeneity problem is created by the potentially simultaneous determination of acceptance wages and jobless spell length. The complete model can be written as

$$\log(w_a) = \mathbf{z_a}\,\boldsymbol{\gamma_a} + \alpha_a \log t + \beta_a \log(w_p) + u_a \qquad (3)$$

$$\log(t) = \mathbf{z}\,\boldsymbol{\gamma_d} + \beta_d \log(w_p) + u_d \qquad (4)$$

$$emp = 1 \text{ if } (\mathbf{z_e}\,\boldsymbol{\gamma_{1e}} + \beta_e \log(w_p) + u_e > 0) \qquad (5)$$

The first equation is the structural equation of the acceptance wage at re-entry (2); the second is the linear projection of the endogenous jobless spell length; the third equation is the selection equation. The latter yields the inverse Mill's ratio. Thus, the variables $\log(w_p)$, $\mathbf{z}$, $\mathbf{z}_a$ and $\mathbf{z}_e$ represent the exogenous variables, while $\log(t)$ is the endogenous variable. γ_a, α_a, β_a, γ_d, β_d, γ_{1e}, β_e are the parameters to be estimated.

The vector $\mathbf{z}_a$ includes dummies reflecting changes in occupation, sector, and working area, year dummies, log unemployment rates and the inverse Mill's ratio. It does not include attributes of new jobs as controls as they may be endogenous. The variable *emp* is a binary variable indicating employment status and the vector $\mathbf{z}_e$ includes variables for previous job attributes.

[10] In particular, we use the White (1980) heteroskedasticity consistent estimator of the variance-covariance matrix.

We also assume u_e distributed as a N(0,1), and orthogonality between the error terms and the variables included in the vector $\mathbf{z}_a$.

Identification of the structural equations (3) and (4) requires exclusion restrictions, namely at least one variable (instrument) correlated with the causal variable of interest, the spell length log(t), but uncorrelated with any other determinants of the dependent variable.

The reform of the training-at-work contract legislation of 1994 provides valid instruments. The training-at-work contract (*Contratto di Formazione Lavoro*, CFL) started operating in 1985 to enhance youth employment. The program granted employers willing to hire eligible workers a substantial labor cost rebate consisting in a 50% reduction of social security contributions and automatic termination at the end of two years. The program featured also an off-the job training component. At the beginning, eligible people were workers aged 16-29. Several reforms of the program took place over the years. The main one, for our purpose, took place in 1994. Firstly, the age eligibility rule was extended from 16-29 to 16-32. Secondly, employers were allowed to hire new training-at-work workers during year t, only if at least 60% of the CFL workers whose contract terminated in t-1 and t-2 were retained on a permanent basis. Thus, the 1994 reform clearly affected the duration of youth unemployment with no impact on post-jobless wages. We, therefore, use two instruments (included in the vector **z**) : (i) a dummy equal to one if individuals are aged 30-32 in the years 1994-1998 (and zero otherwise) to capture the extension of the workers eligibility; and, (ii) a dummy activated in years 1994-1998 to capture the introduction of the restrictive rule on new CFL hires. Finally, the vector **z** also includes variables referring to previous job attributes (i.e. occupation, type of contract, sector, working area, year of separation, employment duration) following the suggestion by Kiefer and Neumann (1979) and Hui (1991) according to which past job experiences affect the distribution and arrival rate of job offers (and, thus, the jobless duration). These variables are not statistically significant in explaining the post-jobless wages, once other controls are added. Therefore we exclude them from structural equation (3) and use them instead as additional identifying instruments in equation (4). Tests of instruments validity are presented in the Tables APP_2 and APP_3 (Appendix).

In the above model we do not account for the possible effects of unobserved job match or individual heterogeneity. Unfortunately, we cannot use standard methods to control for individual heterogeneity without considering multiple spells. Thus, we follow the method proposed by Topel (1986) and used by many others since. The idea is simple: pre-displacement wages paid to the workers depend on their productivity and their (observable and unobservable) characteristics. Thus, conditioning the post-jobless wage equation on pre-displacement wages (by constraining the coefficient on the pre-jobless wage to be equal to one), yields control for unobservable worker characteristics as education, individual productivity and ability (Bartel and Borjas, 1981).

4. Empirical results

4.1 Smoothed hazard estimates and cumulative re-employment rates by groups

Figure 1 displays the unconditional smoothed estimate of re-employment hazard from the pooled data. The monthly re-employment hazard estimate increases over the first 9 months and then decreases. This pattern may hide the existence of different groups of individuals: one of "quickly re-employed", with increasing hazard estimate, and the other of "slowly re-employed" individuals..

Table 2 shows estimates of the cumulative proportion of re-employed young males, with breakdowns by groups. Periods out of work in early careers are indeed very long. Although one fourth of the unemployed was re-employed after 6 months of joblessness, 40% had yet to find a new position three years after entry in the jobless state. This share was down to 33% for those aged 16-19 on their first job, and reached 51% for the late starters (over 26). On the same score, not surprisingly, individuals with a previous relatively long work experience (12-23 months) were doing better than those with shorter tenure (9-12 months): 33% vs. 48% still unemployed. As already pointed out, these numbers imply longer unemployment duration than reported anywhere else in Western Europe.

Differences in cumulative re-employment rates are found between individuals working in different geographical areas. Young workers of Northern Italy have the highest re-employment rates: after three years, 72% have been re-hired, whereas the corresponding rate for their Southern colleagues is 52%. This is in line with the literature on Italy's regional differences, widely documented by many well known contributions.

The proportion of people remaining jobless is higher for individuals who experience a separation during the 1993 recession.[11] Some of the differences conceal composition effects: marked differences are observed across job sectors, occupations and working contracts: after a three-year jobless spell, only 58% of individuals with permanent contract are re-employed, against 68% of the individuals with training-at-work contracts, and 67% of the trainees. Note that, as will be discussed shortly, long unemployment duration does not necessarily imply negative duration dependence.

4.2 Transition out of joblessness: logistic hazard regression model estimates

In this section, we investigate the factors that impact on the speed of transition from joblessness to employment. Fig. 1 shows the baseline hazard function over the study period (36 months), calculated without controlling for individual heterogeneity. [12] Fig.2 displays the estimated Kaplan-Meier survivor function.

Table 3 yields the coefficients of a log(time) baseline hazard function $\gamma(t) = (q-1)\ t$; the estimates of (q-1) are presented for the pooled sample and for the sample of workers becoming unemployed in period t (with t=1989 to 1995), without taking into account the impact of unobserved heterogeneity. In table 4, we present estimates of the q1,… q6 coefficients of a non-parametric flexible specification (with duration-interval dummies) of the alternative baseline function

$$\gamma(t) = q1 * month(1\text{-}6) + q2 * month(7\text{-}12) + \ldots\ldots q6 * month\ (31\text{-}36) + covariates$$

[11] Only few individuals suffered involuntary separation in 1989. For this reason we do not present cumulative re-employment rates for displacements occurred in 1989.

[12] A similar Kaplan-Meier baseline hazard is reported in France by A. Terracol (2009): individuals who receive guaranteed minimum income (RMI) are compared with people who do not. As expected, the baseline hazard of the latter slightly dominates the former (i.e. it is a bit lower): after 12 months the surviving rate is around 50% (62% in our exploration); after 30 months it is around 42% (25% in our data).

with unobserved heterogeneity accounted for (complete results in tab. 5). While, as expected, the estimates yield -q1^ > -q2^ > …..> -q6^, both specifications suggest that negative unemployment duration dependence is modest compared to what might have been expected given the length of unemployment duration **(Box A).** The reference group includes individuals aged 20-25 years on entry, blue collar workers who had permanent contracts in the manufacturing sector of the Northwest, and experienced separations in 1989.

There are differences in re-employment probabilities associated with age, previous job occupation and previous job sectors. The oldest individuals (aged 26 or more) have lower re-employment probabilities than the reference group. Individuals having training-to-work contracts have a higher re-employment probability. Also individuals with a previous occupation as trainees ("apprendisti") have a higher probability of exiting unemployment. Instead, individuals with jobs in the service industry (except the financial sector) have lower probabilities to be re-employed than in manufacturing (reference group). Longer job experience leads to higher re-employment probability. The above results confirm those reported in table 2 displaying the cumulative re-employment rates. As expected, the estimated elasticity of the hazard estimates with respect to the local unemployment rate is statistically significant and negative (about –0.5): re-employment is lower if job availability is lower. Local unemployment rates are highly correlated with the working areas, and, thus, we include in the regression either the local unemployment rates (model A) or the area dummies (model B). In both specifications we find almost identical results. In particular, individuals working in Southern Italy or the Islands (both high unemployment areas) have the lowest probability of re-employment.

Table 3. Estimated duration dependence: log(time) baseline hazard function

Regression for each displacement year:	Estimated duration dependence	Unobserved heterogeneity	Covariates
1989	-0.381* (0.189)	no	Yes
1990	-0.532** (0.057)	no	Yes
1991	-0.713** (0.052)	no	Yes
1992	-0.543** (0.055)	no	Yes
1993	-0.440** (0.064)	no	Yes
1994	-0.439** (0.080)	no	Yes
1995	-0.498** (0.159)	no	Yes
Pooling sample	-0.567** (0.026)	no	Yes

Note: the covariates are the same variables used in table 5

Table 4. Estimated duration dependence: non-parametric baseline hazard function

Pooling sample	Estimated duration dependence	Estimated duration dependence
Unobserved heterog	no	yes
Covariates	yes	yes
months 1-6	-4.048** (0.840)	-4.189** (0.970)
months 7-12	-4.583** (0.841)	-4.621** (0.971)
months 13-18	-4.671** (0.841)	-4.637** (0.970)
months 19-24	-5.257** (0.844)	-5.171** (0.972)
months 25-30	-5.229** (0.843)	-5.117** (0.971)
months 31-36	-5.736** (0.848)	-5.595** (0.976)

Note: the covariates are the same variables used in table 5

Box A : how big is negative unemployment dependence ?

Equation (1) yields the probability of person i of being re-employed after t months, given that he has been jobless for t-1 months. In order to have orders of magnitude of negative unemployment dependence, we report rough bounds of such probabilities for the benchmark individual, calculated from the coefficient estimates of tab. 3 and 4. For tab. 3 we use a linear approximation of $\gamma(t)$ obtained from fig.2 and the estimated coefficient of duration dependence on the pooled sample ; for tab. 4 we make direct use of the (q)-coefficients applied to the duration interval dummies.

Pr[re-employed after T months| unemployed through T-1]	**Tab. 3**		**Tab. 4 Coeff – 1 sigma**	**Tab.4 Coeff.**	**Tab.4 Coeff. + 1 sigma**
T = 3.5 (mid-interval 1-6 months)	**0.015**		0.101	**0.23**	0.453
T = 15.5 (mid-interval 13-18 months)	**0.010**			**0.16**	
T = 27.5 (mid-interval 25-30 months)	**0.006**		0.042	**0.11**	0.237
Negative unemployment duration dependence as shown by the difference of conditional probabilities: Pr[re-employed after T months| unemployed til T-1] - Pr[re-employed after V months| unemployed til V-1]					
T = 3.5; V = 15.5	**0.005**			**0.07**	
T = 3.5; V = 27.5	**0.009**		0.06	**0.12**	0.21

Negative unemployment duration dependence appears very modest. All estimates of Pr[re-employed after T months| unemployed through T-1], whether obtained from the estimated coefficients reported in Tab. 3 or 4, are small. For a short unemployment spell like T =3.5 the largest estimate of the re-employment probability does not exceed 0.23. After a long unemployment spell (V = 27.5 months = over 2 years) the re-employment probability of 0.11 may seem reasonable. But the difference between the two is 0.12, which is indeed very small to claim negative duration dependence. It means that the probability of re-employment after a long unemployment spell (2 years +) is only twice as large as the same probability after an unemployment spell of 3.5 months, about 12 p.p. higher. A sizeable negative unemployment duration ought to imply a much larger difference between the two. Even if we stretch Tab. 4, adding/subtracting one standard deviation to the estimated coefficients, the difference of 21 p.p. between the two conditional re-employment probabilities is still too small to prove the existence of strong negative duration dependence.

4.3 Simultaneous estimation of post-joblessness acceptance wage and elapsed job duration

Post-joblessness acceptance wage and elapsed job duration are jointly determined and simultaneously estimated by 3SLS [13]. Identification requires exclusion restrictions: two instruments have been identified in par. 3.2, their interpretation being discussed while presenting the results of the equation explaining jobless duration. Tab. 6 and 7 display the complete results.

Two specifications have been estimated: in specification 1 the coefficient of elapsed jobless duration is unconstrained, while in it is constrained to 1 in specification 2 (as in Topel, 1989). Both versions deliver very similar results, aside from the constrained coefficient. Consider first the equation that explains post-joblessness acceptance wage (tab. 6): an increase in jobless duration of 10% will lower wages on the subsequent job by 0.4% This implies, for instance, that the doubling of jobless spell duration (on average, from 6 to 12 months) will reduce the wage at re-employment by a slight 4%: for the average individual, it means a wage reduction from 1565 euro to 1502 euro. Declining reservation wages / human capital depreciation/ stigma effects may be associated with longer jobless duration, but the Italian case appears to be different from other cases. Our estimated wage loss incurred by young males is much smaller than the 10-15% estimated on the total working population in Unites States, Canada, UK and France (Ruhm, 1991; Addison and Portugal, 1989; Topel, 1990; Jacobson et al, 1993; Farber, 1993 and 1997; Seninger, 1997; Houle and Van Audenrode, 1995 Gregory and Jukes, 1997; Cohen, Lefranc and Saint-Paul, 1997). It is more, in line with results reported by Rosolia and Saint-Paul (1998) and Kletzer and Fairlie (2003). A plausible, underlying reason is the existence in Italy of an informal wage floor below which people will not be hired: in fact, Social Security and health insurance benefits are granted provided the employee's pay is above a minimum which, in the late Nineties, was equivalent to about 50 eu / day.

The estimated elasticity of pre-jobless wages (when freely determined) is positive and about 3.7%. Notable findings are the positive effects of changes in working area (about +10

[13] Post-joblessness acceptance wages have been separately estimated also by OLS with and without selectivity adjustment. Results and comments on interesting differences with the simultaneous version are in Appendix.

%) and occupational change (from 15% to 24%), due to a switch to higher wage areas and higher qualification. Sector changes have no significant effects on wages as in US studies, showing minor wage losses for workers who switch industry. Also local regional unemployment rates have no direct statistical significant effects. The estimated coefficient of the inverse Mill's ratio is positive (in the unconstrained version) and significant suggesting that currently jobless individuals have greater wage losses than their employed counterparts: a signal that both the duration and the selectivity arguments are capturing the (modest) impact of declining reservation wage and human capital depreciation on post-jobless wages.

Consider now the equation of elapsed jobless duration (tab. 7), whose predicted values (expressed in log months) feed into the equation of acceptance wages. All regressors are strictly exogenous, and reflect the individual workers' observables (previous earnings, age at displacement, job and industry characteristics, working area, displacement year). In addition two identifying instruments are included, which directly impact on jobless duration, but not on re-employment wages.[14] Both reflect changes in labor legislation: the first is a dummy activated when individuals are aged 30-32 and the calendar year is 1994-1998 (<worker eligibility extension 1994>): it captures the extension of age eligibility for training-and-work (CFL) contracts, relevant for about 3% of the sample observations. The second dummy (<employer eligibility criteria 1994>) is activated in the calendar year is 1994-1998 and captures the introduction of a new employers' eligibility rule, restricting the utilization of training-and-work contracts, and touching upon 18% of the observations. The coefficient of the second instrument is positive and statistically significant indicating that the introduction of the new rule increases the length of unemployment as it restricts the numbers of training-and-work contracts offered by firms. The coefficient of the first dummy has the expected negative sign (although not statistically significant), suggesting that the age extension could reduce unemployment length as a larger number of young workers become eligible for training-and-work contracts.

[14] Validity of the instruments has been tested and displayed in Tab. A_2 and A_3.

4.4 Robustness

We conclude with a robustness test aimed at checking our definition of involuntary separations. As explained in par. 2 analysis must exclude individuals who voluntarily separated from their jobs. So far, involuntary separations have been recognized as those of individuals who experience a jobless period longer than one month. Here we re-estimate the model using an alternative, more stringent definition whereby involuntary separations are those associated to jobless periods longer than three months. The only different result is a slight increase in the impact of jobless duration on re-entry wages: a 10% increase in jobless duration will lower wages on the next job by 0.5% . [15] All the remaining results are confirmed.

5. Conclusions

In this paper, we explore the mechanisms that produce stigma, discouragement, and human capital depreciation in the course of joblessness. In particular, we investigate the existence of negative jobless duration dependence and the impact of jobless spells on future wages. Our findings are, to some extent, surprising and relatively out of line compared to analogous explorations in countries other than Italy.

First of all, we find strong evidence of very long unemployment duration of the young male labor force: in the course of the Nineties 40% of the unemployed had yet to find a job 36 months after entry in the jobless state. This share was down to 33% for those aged 16-19 on their first job, and reached 51% for the late starters (over 26). These numbers imply much longer unemployment duration than reported anywhere else in Western Europe.

In second place, and despite our findings on unemployment duration, we find that negative unemployment duration dependence is modest. While the probability of re-employment decreases also in Italy as the elapsed jobless period becomes longer, such decay is small.

[15] The estimates of the model are available upon request.

Thirdly, we show that young Italian males experiencing jobless periods in their early careers face slight re-employment wage losses. Such losses do increase with the duration of joblessness, but here, once again, they are much lower than reported in Unites States, Canada, UK, France and Spain. Our explanation for the Italian specificity is the existence of a wage floor below which people will not be hired as they would lose the eligibility to Social Security and health insurance benefits: in the late Nineties the wage floor was set at about 50 eu /day. Such floor is not a legal "minimum wage" as it exists in many other countries, but it has similar consequences on all forms of "regular" dependent work. The new flexible contracts introduced by the 1997 legislation (Pacchetto Treu) and excluded from this exploration, are not subject to similar provisions. As explained at the outset, their exclusion does not impair the results of analysis, as their impact on youth employment became substantial only since the late 1998.

A final note of caution: we have restricted attention to young workers characterized by a relatively high attachment to the labor market, reflected by the fact that all the sample individuals have had an initial working spell at least 9 months long. As pointed out at the beginning of our paper, this might introduce a dose of "optimistic" bias in our estimates. Had we extended the analysis to individuals representing the universe of young Italian males, in particular of those who have had short initial working spells, our results would look quite different, possibly more in line with the estimates reported for other countries.[16] This is an important task that we leave for future research.

[16] Preliminary evidence is reported in a yet unpublished script (B. Contini and E. Grand, 2010).

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Figure 1. Smoothed hazard estimate

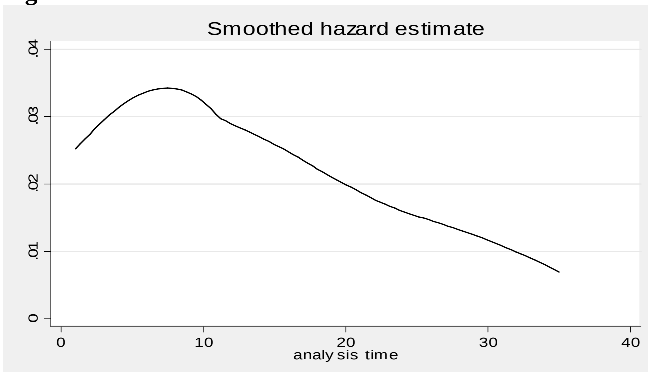


Figure 2. Kaplan-Meier survivor function (survivor in joblessness)

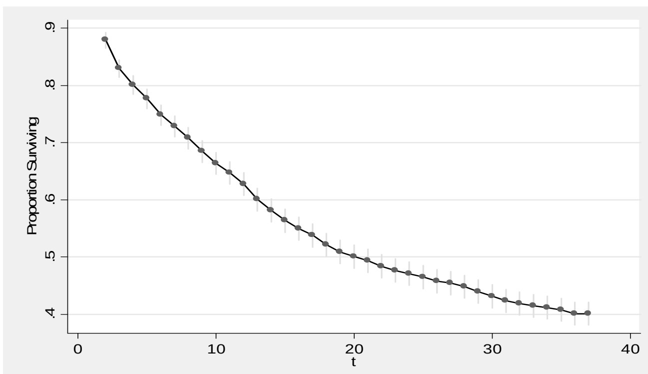


Table 1. Descriptive statistics

Number of observations	2318	Previous job sector	
Working area		Industry	42.84
Northwest	25.33%	Construction	24.81
Northeast	15.32%	wholesale. automotive and repair	17.3
Centre	20.85%	Entertainment	6.21
South	26.07%	transportation. communication	2.29
Islands	12.43%	Finance	0.91
Displacement year		services. research and real estate	5.65
1989	4.31%	Previous job contract	
1990	19.28%	Permanent	79.26%
1991	23.47%	training-at-work	20.74%
1992	21.48%	Previous job occupation	
1993	18.29%	Trainee	31.79%
1994	10.48%	blue collar	53.15%
1995	2.67%	white collar	15.06%
Age at initial period		Previous job experience	
16-19 years	37.62%	9-12 months	45.90%
20-25 years	41.50%	12-23 months	54.10%
26 or more years	20.88%	real monthly wage: mean (std dev)	1318 (449)

Table 2. Cumulative proportion of re-employed individuals

	Months			
	6	12	24	36
All	0.27	0.4	0.53	0.6
Working area				
Northwest	0.33	0.45	0.62	0.68
Northeast	0.4	0.54	0.66	0.72
Centre	0.24	0.38	0.51	0.57
South	0.19	0.32	0.44	0.52
Islands	0.18	0.32	0.45	0.51
Displacement year				
1990	0.3	0.43	0.6	0.66
1991	0.31	0.44	0.57	0.6
1992	0.29	0.4	0.52	0.61
1993	0.21	0.35	0.49	0.56
1994	0.25	0.4	0.54	0.65
1995	0.31	0.5	0.66	0.71
Age at initial period				
16-19 years	0.26	0.41	0.59	0.67
20-25 years	0.28	4	0.52	0.59
26 or more years	0.27	0.37	0.45	0.49
Previous job sector				
Industry	0.3	0.44	0.59	0.66
Construction	0.22	0.36	0.48	0.54
wholesale. automotive and repair	0.26	0.37	0.51	0.57
Entertainment	0.28	0.41	0.54	0.58
transportation. communication	0.26	0.4	0.47	0.47
Finance	0.33	0.48	0.67	0.67
services. research and real estate	0.25	0.37	0.47	0.54
Previous job contract				
Permanent	0.25	0.38	0.52	0.58
training-at-work	0.34	0.46	0.62	0.68
Previous job occupation				
Trainee	0.28	0.42	0.59	0.67
blue collar	0.26	0.38	0.5	0.56
White collar	0.3	0.44	0.55	0.59
Previous job experience				
9-12 months	0.22	0.32	0.46	0.52
12-23 months	0.31	0.44	0.6	0.67

Table 3. Estimated duration dependence: log(time) baseline hazard function

Regression for each Displacement year:	Estimated duration dependence	Unobserved heterogeneity	Covariates
1989	-0.381* (0.189)	no	Yes
1990	-0.532** (0.057)	no	Yes
1991	-0.713** (0.052)	no	Yes
1992	-0.543** (0.055)	no	Yes
1993	-0.440** (0.064)	no	Yes
1994	-0.439** (0.080)	no	Yes
1995	-0.498** (0.159)	no	Yes
Pooling sample	-0.567** (0.026)	no	Yes

Note: the covariates are the same variables used in table 5

Table 4. Estimated duration dependence: non-parametric baseline hazard function

Pooling sample	Estimated duration dependence	Estimated duration dependence
Unobserved heterog	no	yes
Covariates	yes	yes
months 1-6	-4.048** (0.840)	-4.189** (0.970)
months 7-12	-4.583** (0.841)	-4.621** (0.971)
months 13-18	-4.671** (0.841)	-4.637** (0.970)
months 19-24	-5.257** (0.844)	-5.171** (0.972)
months 25-30	-5.229** (0.843)	-5.117** (0.971)
months 31-36	-5.736** (0.848)	-5.595** (0.976)

Note: the covariates are the same variables used in table 5

Table 5. Transition from joblessness to employment (non-parametric baseline hazard function)

	Model A			Model B		
Variables	Coef.		Std.Err.	Coef.		Std.Err.
Age in the displacement year is 16-19	-0.170		0.090	-0.172		0.090
Age in the displacement year is 26 or more	-0.298	**	0.087	-0.307	**	0.087
Previous job occupation: trainees	0.345	**	0.104	0.350	**	0.104
Previous job occupation: white collars	0.095		0.104	0.106		0.104
Previous job contract: training-at-work	0.337	**	0.087	0.348	**	0.087
Previous job sector: construction	-0.209	*	0.085	-0.204	*	0.085
Previous job sector: wholesale, automotive, repair	-0.222	*	0.095	-0.211	*	0.095
Previous job sector: entertainment	-0.083		0.140	-0.051		0.142
Previous job sector: transportation, communication	-0.173		0.239	-0.164		0.240
Previous job sector: finance	0.157		0.344	0.145		0.345
Previous job sector: services, research, real estate	-0.364	*	0.157	-0.361	*	0.158
Previous job experience: ln(months)	0.495	**	0.115	0.488	**	0.115
Previous job earnings: log(real monthly wage)	0.002		0.123	-0.001		0.124
log(regional unemployment rate)	-0.505	**	0.064	---		---
Previous working area: Northeast	---		---	0.211	*	0.100
Previous working area: Centre	---		---	-0.331	**	0.096
Previous working area: South	---		---	-0.511	**	0.092
Previous working area: Islands	---		---	-0.530	**	0.118
Displacement year is: 1990	1.255	**	0.226	1.271	**	0.226
Displacement year is: 1991	1.147	**	0.226	1.145	**	0.226
Displacement year is: 1992	1.110	**	0.226	1.097	**	0.226
Displacement year is: 1993	0.965	**	0.230	0.956	**	0.230
Displacement year is:1994	1.130	**	0.240	1.090	**	0.240
Displacement year is: 1995	1.363	**	0.292	1.323	**	0.293
Baseline hazard function: month 1-6	-4.189	**	0.970	-5.090	**	0.965
Baseline hazard function: month 7-12	-4.621	**	0.971	-5.526	**	0.966
Baseline hazard function: month 13-18	-4.637	**	0.970	-5.548	**	0.965
Baseline hazard function: month 19-24	-5.171	**	0.972	-6.088	**	0.967
Baseline hazard function: month 25-30	-5.117	**	0.971	-6.038	**	0.966
Baseline hazard function: month 31-36	-5.595	**	0.976	-6.517	**	0.971
sigma_u	0.655	**	0.138	0.655		0.139
Rho	0.115	**	0.024	0.115		0.024
log-likelihood	-5820.07			-5816.77		

Note: the reference group is composed by blue collars aged 20-25 years old (that suffer involuntary separations in 1989), working with standard contract in manufacture in the Northwest; ** means statistical significant at 1% level; * means statistical significant at 5% level.

Table 6. The determinants of post-joblessness wages

log(acceptance wage)	3SLS			3SLS		
	Specification 1			Specification 2		
	Coef.		Std. Err.	Coef.		Std. Err.
Elapsed jobless duration: log(months)	-0.039	***	0.008	-0.042	***	0.010
Previous job earnings: log(real monthly wages)	0.371	***	0.030	1.000		.
Dummy: change in working area	0.130	***	0.030	0.096	**	0.038
Dummy: sector change	-0.007		0.018	-0.027		0.022
Dummy: occupational change	0.149	***	0.020	0.241	***	0.027
Log (regional unemployment rate)	-0.010		0.019	0.007		0.023
Lambda	0.167	***	0.053	0.007		0.060
Year dummies	yes		yes	yes		Yes
_cons	4.308	***	0.217	-0.108		0.351
R-squared	0.1990			0.1961		

Note: *** statistical significant at 1% level ; ** at 5% level; * at 10% level;

Table 7 -. Elapsed jobless duration: first step estimation and selection equation

	First step estimation (specification 1)			**First step estimation (specification 2)**			**Selection equation**		
	Elapsed jobless duration: log(months)			Elapsed jobless duration: log(months)			Emp		
	Coef.		S. E.	Coef.		S. E.	Coef.		S. E.
Previous job earnings: log(real monthly wages)	0.079		0.135	0.152		0.135	0.091		0.105
age in the displacement year	0.054		0.045	0.055		0.045	-0.055	***	0.010
Previous job occupation: trainees	0.046		0.123	0.041		0.123	0.084		0.091
Previous job occupation: white collars	-0.327	***	0.109	-0.317	***	0.109	0.060		0.090
Previous job contract: training-at-work	-0.273		0.217	-0.273		0.217	0.259	***	0.077
Previous job sector: construction	0.099		0.162	0.098		0.162	-0.180	**	0.074
Previous job sector: wholesale, automotive, repair	0.25	*	0.14	0.249	*	0.140	-0.145	*	0.082
Previous job sector: entertainment	-0.057		0.129	-0.063		0.129	-0.039		0.120
Previous job sector: trasportation, communication	-0.21		0.246	-0.214		0.246	-0.152		0.189
Previous job sector: finance	0.195		0.336	0.200		0.337	0.188		0.307
Previous job sector: services, research, real estate	0.273		0.277	0.274		0.278	-0.289	**	0.131
Previous working area: Northeast	-0.361	**	0.141	-0.362	**	0.141	0.157	*	0.094
Previous working area: Centre	0.278		0.2	0.281		0.201	-0.243	***	0.084
Previous working area: South	0.493	**	0.253	0.499	**	0.253	-0.305	***	0.080
Previous working area: Islands	0.534	*	0.314	0.538	*	0.314	-0.379	***	0.098
Displacement year: 1990	-1.163		0.852	-1.178		0.853	0.904	***	0.155
Displacement year: 1991	-1.193		0.726	-1.208	*	0.726	0.722	***	0.154
Displacement year: 1992	-1.084		0.763	-1.096		0.764	0.775	***	0.154
Displacement year: 1993	-1.151	*	0.673	-1.163	*	0.674	0.654	***	0.157
Displacement year: 1994	-2.117	***	0.808	-2.136	***	0.809	0.839	***	0.171
Displacement year: 1995	-2.981	***	0.839	-2.991	***	0.839	0.855	***	0.232
Tenure: log(months)	-0.462		0.275	-0.459	*	0.275	0.336	***	0.101
Worker eligibility extension 1994	-0.34		0.245	-0.348		0.245			
Employer eligibility criteria 1994	1.699	***	0.116	1.704	***	0.116			
Lambda	-1.645		1.383	-1.687		1.384			
_cons	3.185		1.994	2.678		1.995	-0.696		0.810

Note: the reference group is composed by blue collars that suffer involuntary separations in 1989, working with standard contract in manufacture in the Northwest; *** means statistical significant at 1% level; ** means statistical significant at 5% level; * means statistical significant at 10% level;

APPENDIX 1

OLS estimation of the wage equation

The Mincerian wage regression is also estimated by ordinary least squares (with and without selectivity adjustment).

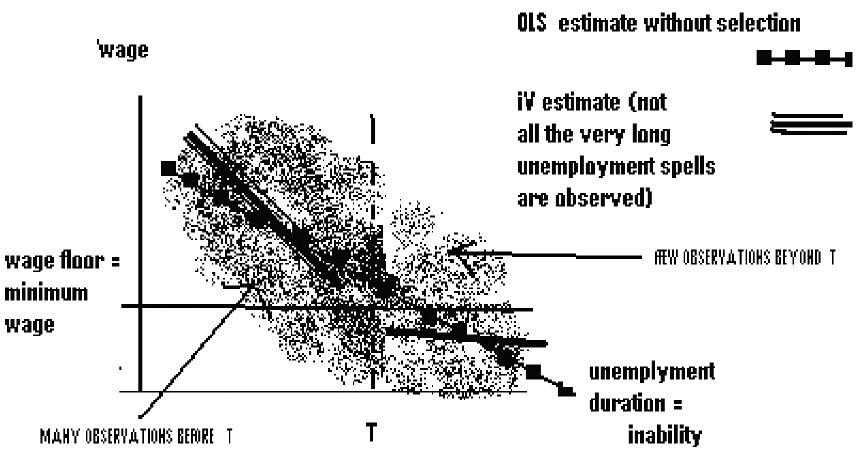
Table A_1 : OLS estimation of the determinants of post-joblessness wages

log(acceptance wage)	OLS			OLS + Heckman		
	Coef.		Std. Err.	Coef.		Std. Err.
Elapsed jobless duration: log(months)	-0.012		0.008	-0.014	*	0.008
Previous job earnings: log(real monthly wages)	0.383	***	0.026	0.371	***	0.027
Dummy: change in working area	0.139	***	0.032	0.126	***	0.032
Dummy: sector change	0.005		0.019	-0.007		0.019
Dummy: occupational change	0.135	***	0.020	0.145	***	0.020
Log (regional unemployment rate)	0.019		0.017	-0.014		0.020
Lambda	no		No	0.157	***	0.050
Year dummies	yes		Yes	yes		yes
_cons	4.303	***	0.357	4.288	***	0.353
R-squared	0.1886			0.1952		

Note: *** means statistical significant at 1% level ; ** means statistical significant at 5% level; * means statistical significant at 10% level;

Not surprisingly, the noticeable difference with the results of simultaneous estimation is the impact of elapsed jobless duration: here an increase in jobless duration of 10% is estimated to lower future wages by a mere 0.12% (0.14% with selectivity adjustment). All the other coefficient estimates are very similar to those obtained with simultaneous estimation.

The fact that the 3SLS estimate is larger (in absolute value) than the OLS estimate may appear counterintuitive if workers exiting jobless spells later are negatively selected based on ability. The explanation runs as follows: we know that the number of re-entering workers who have experienced a very long unemployment spell is relatively small, given that our observation window ends in 1998: between ¼ and 1/3 of all the entrants from 1992 onwards will not be observed if the unemployment spell exceeds 3 years (table 2). Thus, we have a problem of selection due to end-of-period truncation. Let T be the threshold (a region, not necessarily a point) beyond which few observations are at hand. It is reasonable to assume that the wage reduction as a consequence of joblessness will reach a negotiated floor below which few will go (the minimum wage in countries other than Italy). The wage schedule as a function of joblessness will therefore be negatively sloped and upwards concave, becoming horizontal approaching the floor (or kinked near T) . Simultaneous estimation of our log-specification, accounting for selectivity, will therefore mainly catch the observations placed to the left of T, while the OLS estimate will reflect the full sample composition. This being the case, the coefficient of the elapsed jobless duration will be more negative under simultaneous estimation than in the OLS version.



APPENDIX 2

The validity of the two instruments corresponding to policy changes used to identify eqs. (3) and (4) is tested by estimating the post-jobless wage equation (specification 1) with their inclusion among the regressors. Table APP_2 shows that the null hypothesis stating that both coefficients are zero cannot be rejected. In Table APP_3 a similar test is performed on a more stringent null hypothesis stating that all the previous job attributes have zero coefficients in the post-jobless acceptance wage. Here too, the null of all coefficients being equal to zero cannot be rejected, suggesting the validity of such attributes as potential instruments.

	Table A_2 = Test for validity of instruments (two policy changes)		
Ho:			
1	[log(acceptance wage)] Worker eligibility extension 1994	=	0
2	[log(acceptance wage)] Employer eligibility criteria 1994	=	0
	chi2(2) = 3.30		
	Prob > chi2 = 0.1918		
	Table A_3 = Test for validity of instruments (previous job attributes)		
Ho:			
1	[log(acceptance wage)] Worker eligibility extension 1994	=	0
2	[log(acceptance wage)] Employer eligibility criteria 1994	=	0

3	[log(acceptance wage)] Previous job occupation: trainees	=	0
4	[log(acceptance wage)] Previous job occupation: white collars	=	0
5	[log(acceptance wage)] Previous job contract: training-at-work	=	0
6	[log(acceptance wage)] Previous job sector: construction	=	0
7	[log(acceptance wage)] Previous job sector: wholesale, automotive, repair	=	0
8	[log(acceptance wage)] Previous job sector: entertainment	=	0
9	log(acceptance wage)] Previous job sector: trasportation, communication	=	0
10	[log(acceptance wage)] Previous job sector: finance	=	0
11	[log(acceptance wage)] Previous job sector: services, research, real estate	=	0
12	[log(acceptance wage)] Previous working area: Northeast	=	0
13	[log(acceptance wage)] Previous working area: Centre	=	0
14	[log(acceptance wage)] Previous working area: South	=	0
15	[log(acceptance wage)] Previous working area: Islands	=	0
16	[log(acceptance wage)] Displacement year: 1990	=	0
17	[log(acceptance wage)] Displacement year: 1991	=	0
18	[log(acceptance wage)] Displacement year: 1992	=	0
19	[log(acceptance wage)] Displacement year: 1993	=	0
20	[log(acceptance wage)] Displacement year: 1994	=	0
21	[log(acceptance wage)] Displacement year: 1995	=	0
22	[log(acceptance wage)] Tenure: log(months)	=	0
	chi2(22) = 22.95		
	Prob > chi2 = 0.1923		